# Inflation Bias after the Euro: Evidence from the UK and Italy

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# Inflation Bias after the Euro: Evidence from the UK and Italy

by

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**Abstract**

This paper presents an indirect approach to investigate the possible existence of measurement error bias in the Harmonized Index of Consumer Prices for the UK and Italy. Our empirical results show that there is no significant evidence of a bias for the UK, nor for Italy prior to the introduction of the Euro. Since January 2002, however, the inflation rate in Italy has been underestimated by at least 6 percentage points.

**Keywords**: Inflation; Measurement bias

**JEL Classification**: C22, C43, D14, E31

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## 1. Introduction

The inflationary consequences of the adoption of the common currency have arguably been one of the most debated issues in recent years in Europe, both in the academic literature and in the policy debate. There is widespread concern over the effect of the currency changeover on price inflation and its negative consequences for the purchasing power of households (see, for example, European Central Bank [March 2002], [April 2002] and [July 2002]). It has been argued that the switch from national currencies to the Euro has served as a device enabling non-competitive firms to co-ordinate upwards their expectations on pricing behaviour. The exogenous change in cash denomination has thus shifted industries with market power to a higher-price equilibrium. In other words, the possibility of permanent price increases in these industries associated with the introduction of the new currency has generated self-fulfilling inflationary expectations and therefore discontinuous price jumps totally unrelated to underlying market conditions or fundamentals (see Adriani, Marini and Scaramozzino, 2003, for an interpretation along these lines).

However, the perception of generalized and permanent price increases triggered by the introduction of the Euro is not supported by the official statistics. Figure 1 describes the evolution of the inflation rate in the euro area and in two leading countries, the UK and Italy, during the past five years. No systematic increases in inflation would appear to have occurred since January 2002, contrary to public perception in Euro countries.

This paper investigates the extent of possible measurement errors in the Harmonized Index of Consumer Prices (HICP) in Italy and the UK following the changeover. Italy has adopted the Euro in January 2002 and has experienced wide jumps in sectoral prices due to allegedly speculative behaviour, whereas the UK has maintained its national currency and has not been affected by the changeover.

The difficulty of accurately measuring inflation in a dynamic economy may be a crucial reason for the decision by central banks not to adopt strict inflation targets (see, *e.g.*, Greenspan, 2001). Following the Boskin Report in the

U.S. (see Boskin, Dulberger, Gordon, Griliches and Jorgenson, 1996)[1], several studies dealing with the evaluation of mismeasurement in national consumer price indexes (CPIs) in European countries have been carried out (see, *e.g.*, Oulton, 1995, Cunningham, 1996, Lequiller, 1997, Hoffman, 1998, Folktersma, 1998, Blow and Crawford, 1999, 2001, and Covas and Silva, 2001). All these studies lend support to the view that consumer price inflation statistics tend to overstate the true inflation rate, as argued by the Boskin Commission in 1996 for the U.S. CPI. This upward bias is believed to reflect both the use of fixed expenditure weights in price indexes (the so-called substitution bias) and quality changes[2]. In particular, Cunningham (1996) estimates that the extent of the bias in the UK Retail Price Index (RPI) is between 0.35 and 0.8 percentage points per year. Blow and Crawford (1999, 2001), using non-parametric methods, estimate that the UK RPI was overstated by up to 3.2 percentage points over the period 1976-1997 relative to the true cost of living, due to the substitution bias.

In the present paper, we infer the overall inflation bias both for the UK and Italy by using an indirect methodology first suggested by Nordhaus (1998). We construct an estimate of the bias on the basis of whether changes in the deflated median income are consistent with survey data on households responses about net changes in their financial situation. Specifically, inflation is correctly measured if, on average, the median real income shows no change when an equal number of families report an improvement and a worsening in their financial condition. Otherwise, there would be a measurement bias. Thus, if more households report a worsening in their financial situation rather than an improvement at a time when the deflated median household income does not change according to official statistics, one would infer that inflation has been underestimated.

An advantage of this methodology is that it can be more robust with respect to the potential misperception of the inflation rate by households than direct surveys on average increases in consumption prices. Since agents directly observe the evolution of their net financial balances at the end of each period, they

[1] The report estimates that the U.S. CPI has an upward bias of 1.1 percentage points per annum.
[2] See, *e.g.*, Abraham, Greenlees and Moulton, 1998, Deaton, 1998, Diewert, 1998, Pollak, 1998, Abraham, 2003, Hausman, 2003, and Schultze, 2003.

are not affected by the relative frequency of purchase that could distort their perception of the inflation rate[3]. Nordhaus (1998), in his seminal paper, shows that the U.S. CPI permanently overstated inflation. According to the report of the Boskin Commission, this result can be accounted for by both the presence of fixed expenditure weights in the index and quality changes, including, for example, the appearance of new goods and improvements in products and in the distribution networks.

Our results are in net contrast with previous findings. There is no statistically significant evidence of a bias in measuring inflation for the UK over the whole period 1996-2003. For Italy, prior to the adoption of the Euro the bias is also not significantly different from zero. After the changeover, however, the rate of inflation has been significantly underestimated. Our findings thus provide empirical support to the view that official statistics failed to capture the upsurge of speculative behaviour, due to implicit coordination of price setters towards high price equilibria, made possible by money illusion and/or market imperfections (Fehr and Tyran, 2001, 2004, and Adriani, Marini and Scaramozzino, 2003).

The structure of the paper is as follows. Section 2 provides a description of the indirect methodologies to estimate measurement errors in price indexes. In Section 3 we describe the data and the empirical specification adopted in the paper. The findings for the UK and Italy are reported in Sections 4 and 5 respectively. Section 6 summarizes the main conclusions.

[3] The alleged gap between perceived and effective inflation is commonly attributed to the fact that consumers would attach greater weight to price changes in goods and services bought more frequently relative to the so-called “big ticket items”, such as durable goods (see, for instance, ECB 2002).

## 2. Indirect approaches to inferring measurement errors in price indexes

The measurement bias in price indexes has recently been analysed by employing indirect methodologies, following the original contribution by Nordhaus (1998). The key idea is to capture the measurement bias by comparing the variation in median income, deflated by the consumer price index, with the households' reported changes in their net financial situation. When the real median income of households increases in a given period of time, a majority of respondents should also report an improvement in their net financial situation. In general, there would be no bias in the measurement of inflation if, on average, the Consumer Price Index (CPI)-deflated median income shows no change when an equal number of individuals report themselves to be better off or worse off. If the deflated median income shows an increase, the price index is affected by a negative bias and inflation is underestimated. By contrast, if the deflated real income experiences a decline, the inflation rate is overestimated. The econometric specification adopted by Nordhaus takes the following form:

$$NIFS_t = \beta_1 \left( \Delta RMed_t - \beta_0 \right) + u_t , \qquad (1)$$

where $NIFS_t$ denotes the percentage of surveyed households reporting a net improvement in their financial situation minus the percentage of households reporting a net worsening in their financial situation over the previous 12 months, $\Delta RMed_t$ is the growth rate of the deflated median household income over the previous year and $u_t$ is a stochastic disturbance. There is no bias if the estimation of the parameter $\beta_0$ is not significantly different from zero. By contrast, if the estimated parameter $\beta_0$ is significantly positive (negative), the consumer price index understates (overstates) inflation. Specifically, $\beta_0$ can be interpreted as an estimate of the bias in the consumer price index, since it corresponds to the percentage variation in measured real income that is associated with an equal number of households reporting an improvement and worsening in their net financial condition.

Using data from the University of Michigan Household Survey of consumer behaviour, Nordhaus (1998) estimates equation (1) to find that the US CPI has been affected by an upward bias of about 1.5 percentage points during the period 1968-1994. Such a result is qualitatively consistent both with the Boskin Commission's estimates and with other literature estimating the bias in inflation by direct measures (see, for instance, Cunningham, 1996, Lequiller, 1997, Hoffman, 1998, Folktersma, 1998, Blow and Crawford, 1999, 2001).

Krueger and Siskind (1998) investigate the issue of the measurement bias using an indirect approach which takes into account changes in the shape of the income distribution. Using data from the Panel Study of Income Dynamics they find that there is no bias in the US CPI over the sample 1968-1991.

In the present paper we extend the indirect approach of Nordhaus by allowing for inertia in the relationship between the net proportion of households that report an improvement in their financial welfare and changes in real median income. In particular, in our empirical analysis we use a dynamic Error Correction Model, which enables us to deal with gradual adjustment in households responses (see Section 3.3).

## 3. Empirical analysis

### 3.1. UK data

We use the UK Government's preferred definition of median disposable income, where housing costs are deducted. This implies that an increase in housing costs would represent a decline in living standards, if the accommodation is unchanged (Goodman and Shephard, 2002).

We utilize data for the households self-assessed financial condition from the European Commission's monthly Harmonized Consumer Survey. Specifically, the relevant question is the following:

> How does the financial situation of your household changed over the last 12 months? [got a lot better; got a little better; stayed the same; got a little worse; got a lot worse; don't know].

Denote by *BB*, *B*, *S*, *WW*, *W*, *D* the percentage of respondents who choose respectively "got a lot better", "got a little better", "stayed the same", "got a little worse", "got a lot worse" and "don't know" (with $BB+B+S+WW+W+D=100$). The aggregate balance, *NIFS*, measures the difference (in percentage points of total answers) between the numbers of people reporting a net improvement in their financial situation and the numbers of people reporting a net worsening in their financial situation, and is calculated as (European Commission, 2003):

$$NIFS = \left(BB + \frac{1}{2}B\right) - \left(WW + \frac{1}{2}W\right) \qquad (2)$$

Let $\Delta RMed$ denote the change in the median income deflated by the HICP. Figure 2 shows that the cyclical pattern of *NIFS* and $\Delta RMed$ tended to be quite similar, in particular since the introduction of euro notes and coins in January 2002. Over the whole sample period from January 2003, *NIFS* displayed an upward trend whereas $\Delta RMed$ did not.

Tables 1a and 1b illustrate formal unit root tests on *NIFS* and $\Delta RMed$. From Table 1a, we cannot reject the hypothesis that *NIFS* is I(1) over the period ending in December 2001. The Akaike Information Criterion (AIC), the Schwarz Bayesian Criterion (SBC) and the Hannan-Quinn Criterion (HQC) for a Dickey-Fuller regression with intercept and a linear trend all support an ADF(1) specification. The value of the test statistic for the ADF(1) is –3.2486, which is greater than the 95% critical value for the augmented Dickey-Fuller statistic (–3.4688). Similar results would hold for the sample period ending in March

2003. Hence there appears to be an element of hysteresis in the changes in the net improvement in the households' financial situation.

The unit root tests in Table 1b are consistent with the hypothesis that $\Delta RMed$ is I(0) over the period ending in December 2001. The AIC, the SBC and the HQC support a DF specification. The value of the test statistic for the DF without linear trend is –7.8816, which is less than the 95% critical value for the augmented Dickey-Fuller statistic (–2.9069) thereby allowing us to reject the null hypothesis of a unit root.

Following the unit root analysis, equation (1) is re-specified to achieve consistency in the order of integration between net improvements in the financial situation and changes in real median income:

$$\Delta NIFS_t = \beta_1\left(\Delta RMed_t - \beta_0\right) + u_t\,, \tag{3}$$

Equation (3) will form the basis for the empirical estimation of the model.

## 3.2. Italian data

We use the measure of median disposable income reported monthly by Istituto di Studi e Analisi Economica (ISAE), which includes labour incomes, capital incomes (including rents and interests) and transfers (pensions) (see Martelli, 2000). Similar to the UK, we use data on the changes in the Italian households financial situation from the European Commission's monthly Harmonized Consumer Survey.

Figure 3 shows that *NIFS* has experienced a sharp decline since January 2002. It is important to note that the stark worsening in the financial situation cannot be attributed to transitory misperceptions regarding the true inflation rate, since the deterioration in the net financial situation becomes more and more severe over time. By contrast, no such decline is apparent in the change in the official real median income, which exhibits marked fluctuations but no downward trend.

The unit root tests in Table 2a cannot reject the hypothesis that *NIFS* is I(1) over the period from 1990 until December 2001. The SBC and the HQC support an ADF(1) specification. The test statistic for the ADF(1) without linear trend is –1.6351, which is greater than the 95% critical value for the augmented Dickey-Fuller statistic (–2.8822). The AIC would tend to lead to a choice of an ADF(2) specification, for which again we cannot reject the I(1) hypothesis. A similar result would hold for the period ending in February 2004.

From Table 2b, $\Delta RMed$ can be regarded as I(0). The AIC, SBC and HQC all support an ADF(1) specification. The test statistic for the ADF(1) without linear trend is –4.3386, less than the 95% critical value for the augmented Dickey-Fuller statistic (–2.8842).

## 3.3. Empirical specification

Our empirical specification of the relationship between the households' survey on financial conditions and changes in real median income is based on the dynamic Error Correction Model (Sargan, 1964; Engle and Granger, 1987). The main advantages of this specification are that it allows for a richer dynamic analysis of the relationship between financial conditions and real median income, and that it is an adequate estimation procedure in the presence of cointegrated variables according to Granger's Representation Theorem.

Since *NIFS* is I(1) and $\Delta RMed$ is I(0), we regress $\Delta NIFS$ on $\Delta RMed$ as in equation (3). The estimated equation is the following:

$$\Delta\Delta NIFS_t = \beta_1 + \beta_2 \Delta\Delta RMed_t + \beta_3 \Delta\Delta NIFS_{t-1} + \beta_4 \Delta\Delta RMed_{t-1} + \beta_5 \Delta NIFS_{t-2} + \beta_6 \Delta RMed_{t-2} + u_t, \quad (4)$$

where $u_t$ is a stochastic disturbance. A non-linear long-run estimator of the bias in reported inflation, measured by HICP, is given by:

$$\hat{\gamma} = -\frac{\hat{\beta}_1}{\hat{\beta}_6}, \tag{5}$$

where $\beta_1$ and $\beta_6$ are the coefficients pertaining to the intercept term and to the change in HICP-deflated median income, respectively. A significantly positive value of $\hat{\gamma}$ would imply that the HICP underestimates the inflation rate and *vice versa*. If $\hat{\gamma}$ is not statistically significant, there is no systematic bias in the estimation of the inflation rate.

## 4. Inflation in the UK

The results of the estimation of the ECM equation (4) for the UK are reported in Table 3. The estimated bias is negative over the pre-2001 period, but it is not statistically significant. There is no residual serial correlation nor heteroskedasticity. The residuals from the regression are I(0), thus confirming that the variables are cointegrated.

An important finding is that there is no significant evidence of a structural break in the relationship after the introduction of the euro in January 2002. The Chow test on the equality of the parameters after 2002 is not statistically significant. When re-estimated over the sample period from June 1996 until March 2003, the inflation bias is still negative and not statistically significant[4].

Hence, there is no evidence of a bias in inflation measurement in the UK, either pre- or post-euro, nor evidence of a structural break following the introduction of the common currency in the euro area.

[4] Direct estimation of equation (3) by non-linear least squares yields similar results.

## 5. Inflation in Italy

Table 4 reports the results of the estimation of equation (4) for Italy for the period from March 1991 to December 2001. The estimated bias is negative but not statistically significant[5]. There is no residual serial correlation nor heteroskedasticity and the regression residuals are I(0), again confirming the cointegration of the series. However, the Chow test reveals a structural break in the relationship after January 2002. This would suggest that differences in median income, deflated by the HICP, no longer predict changes in the households' financial condition consistent with the pre-euro relationship. Our explanation for this structural break is that, after the currency changeover, the price index is no longer an unbiased estimate of the relevant inflation rate.

In order to obtain an estimate of the likely extent of the bias in the price index, we follow an indirect procedure similar in spirit to Fieller's (1954) estimation of confidence intervals. We consider alternative values of the inflation bias and look at the smallest level of the bias for which there is no significant structural break before and after the introduction of the euro. Table 5 illustrates the results of our estimation. We estimate the ECM over the period March 1991 to December 2001 and compute Chow's test for structural stability for the post-euro sample period, from January 2002 to February 2004. With 5% significance level, inflation bias is in excess of 6 percentage points. The significance level of the structural break is higher than 5% only when the inflation bias is no less than 7 percentage points. Hence, the HICP appears to have under-estimated inflation by at least 6 percentage points per year since the introduction of the euro in January 2002.

A further element of robustness of our analysis is the stability of the disposable income distribution in Italy since 1995, which emerges from the ISAE income survey data[6]. Similar findings on Bank of Italy data are reported in Brandolini (2004). Therefore, our results seem to be robust also with respect to

[5] Similar results are obtained by estimation of equation (1) by non-linear least squares.
[6] Details of the analysis of income distribution in Italy are available from the authors upon request.

Krueger and Siskind's (1998) concerns about a time-variant shape of the income distribution.

## 6. Conclusions

This paper has investigated the possible existence of measurement bias in the Harmonized Consumer Price Index for the UK and Italy. There is no evidence of a statistically significant bias for the UK over the period 1996-2003. A similar result has also been found for Italy for the sample ending in December 2001. However, since the introduction of the Euro there is evidence that the inflation rate in Italy has been underestimated. There is an unambiguous structural break in January 2002. Changes in median income, deflated by the official HICP, are inconsistent with the households' reported worsening in their net financial situation.

An economic explanation for this result can be found in generalized speculative behaviour in non-competitive markets. The currency changeover has acted as a device that led firms with market power to increase their prices. The official statistics do not seem to have captured the price jumps in these sectors, and therefore measured inflation does not incorporate these increases.

Our findings show that in Italy there has been a downward bias in measuring inflation since 2002 of at least 6 percentage points. These results are particularly significant since the use of statements on net changes in the financial position of households is immune from the criticism that households might misperceive the inflation rate due to the different relative frequency of purchases of consumption items.

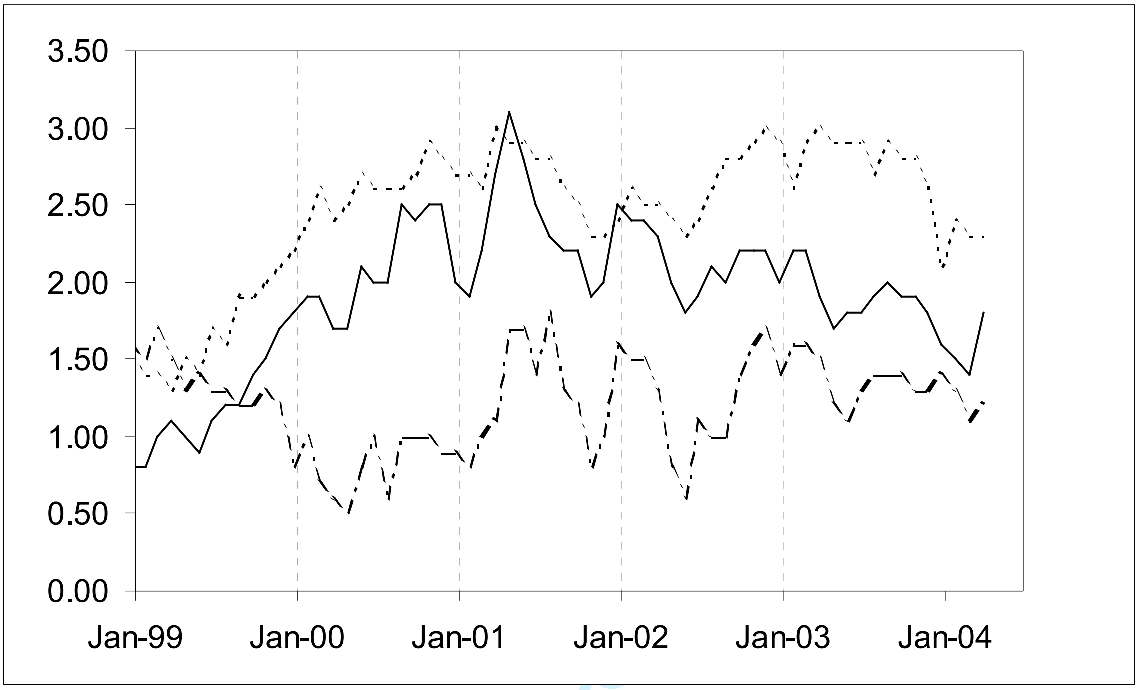
## Data Appendix

| Series | Source |
| --- | --- |
| Harmonized Consumer Price Index UK | Bloomberg |
| Harmonized Consumer Price Index Italy | Bloomberg |
| After Housing Cost (AHC) Median Income UK | IFS |
| Median Disposable Income Italy | ISAE |
| Percentage balance of the answers to the question "How does the financial situation of your household changed over the last 12 months?" UK | European Commission's Business and Consumer Survey Question 1 |
| Percentage balance of the answers to the question "How does the financial situation of your household changed over the last 12 months?" Italy | European Commission's Business and Consumer Survey Question 1 |

Notes:

IFS data were downloaded from: www.ifs.org.uk/inequality/bn19figs.zip.
European Commission data were downloaded from:
http://europa.eu.int/comm/economy_finance/indicators/businessandconsumersur-veys_en.htm.

**Figure 1. Inflation in the Euro Area, the UK and Italy**



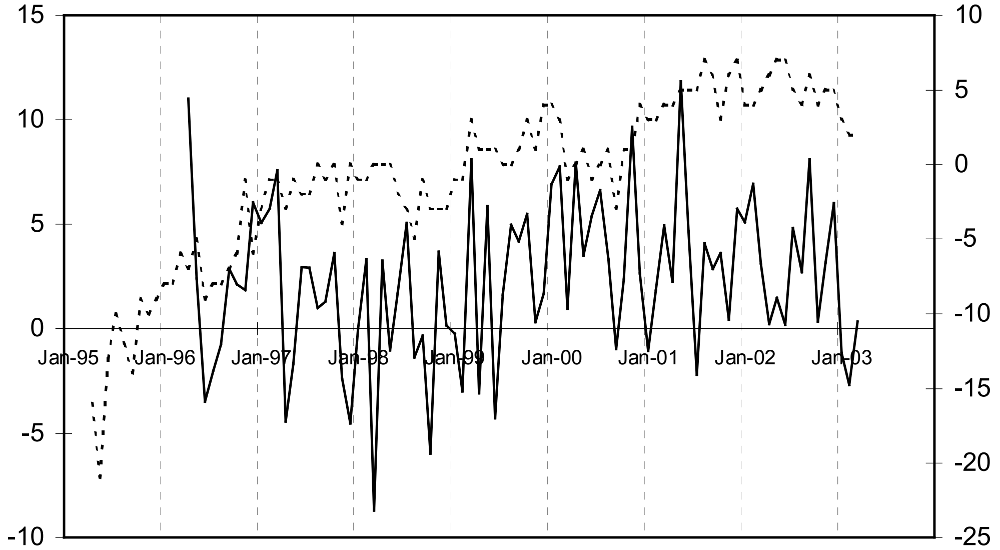

Legenda:

____________ Euro Area

_._._._._._. UK

………….. Italy

**Figure 2. Net improvement in financial situation and changes in real median income in the UK.**

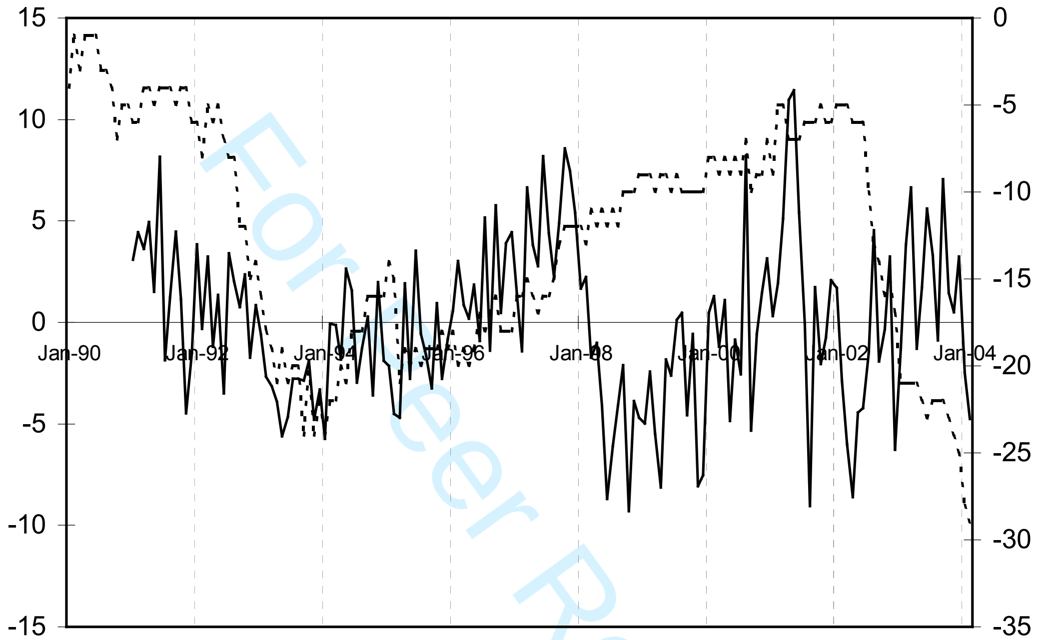



Legenda:

__________ *ΔRMed* (left axis)

………….. *NIFS* (right axis)

**Figure 3. Net improvement in financial situation and changes in real median income in Italy.**




Legenda:

__________ *ΔRMed* (left axis)

………….. *NIFS* (right axis)

**Table 1a. Unit root tests: Net improvement in financial situation, UK.**

Unit root tests for the variable $NIFS_t$

| | Test statistic | LL | AIC | SBC | HQC |
|---|---|---|---|---|---|
| **DF** | –4.8449 | –153.2692 | –156.2692 | –159.7653 | –157.6664 |
| **ADF(1)** | –3.2486 | –150.1057 | –154.1057 | –158.7672 | –155.9687 |
| **ADF(2)** | –2.8221 | –149.5899 | –154.5899 | –160.4167 | –156.9186 |
| **ADF(3)** | –2.6907 | –148.8112 | –154.8112 | –161.8034 | –157.6057 |
| **ADF(4)** | –2.7992 | –147.7837 | –154.7837 | –162.9413 | –158.0438 |

The Dickey-Fuller regressions include an intercept and a linear trend.
Sample period: September 1995 to December 2001.
95% critical value for the augmented Dickey-Fuller statistic = –3.4688.

LL = Maximized log-likelihood
AIC = Akaike Information Criterion
SBC = Schwarz Bayesian Criterion
HQC = Hannan-Quinn Criterion

## Table 1b. Unit root tests: Change in real median income, UK.

Unit root tests for the variable $\Delta RMed_t$

| | Test statistic | LL | AIC | SBC | HQC |
|---|---|---|---|---|---|
| **DF** | –7.8816 | –177.3181 | –179.3181 | –181.4770 | –180.1686 |
| **ADF(1)** | –5.1786 | –177.1161 | –180.1161 | –183.3544 | –181.3918 |
| **ADF(2)** | –4.0659 | –176.8970 | –180.8970 | –185.2147 | –182.5980 |
| **ADF(3)** | –3.1022 | –175.9545 | –180.9545 | –186.3517 | –183.0808 |
| **ADF(4)** | –2.6387 | –175.7651 | –181.7651 | –188.2417 | –184.3166 |

The Dickey-Fuller regressions include an intercept but not a trend.
Sample period: September 1996 to December 2001.
95% critical value for the augmented Dickey-Fuller statistic = –2.9069.

LL = Maximized log-likelihood
AIC = Akaike Information Criterion
SBC = Schwarz Bayesian Criterion
HQC = Hannan-Quinn Criterion

**Table 2a. Unit root tests: Net improvement in financial situation, Italy.**

Unit root tests for the variable $NIFS_t$

| | Test statistic | LL | AIC | SBC | HQC |
|---|---|---|---|---|---|
| **DF** | –1.9689 | –256.3919 | –258.3919 | –261.3264 | –259.5844 |
| **ADF(1)** | –1.6351 | –244.4104 | –247.4104 | –251.8121 | –249.1992 |
| **ADF(2)** | –1.7402 | –242.9674 | –246.9674 | –252.8363 | –249.3524 |
| **ADF(3)** | –1.8048 | –242.3557 | –247.3557 | –254.6919 | –250.3369 |
| **ADF(4)** | –1.9045 | –241.6453 | –247.6453 | –256.4488 | –251.2228 |

The Dickey-Fuller regressions include an intercept but not a trend.
Sample period: June 1990 to December 2001.
95% critical value for the augmented Dickey-Fuller statistic = –2.8822.

LL = Maximized log-likelihood
AIC = Akaike Information Criterion
SBC = Schwarz Bayesian Criterion
HQC = Hannan-Quinn Criterion

**Table 2b. Unit root tests: Change in real median income, Italy.**

Unit root tests for the variable $\Delta RMed_t$

| | Test statistic | LL | AIC | SBC | HQC |
|---|---|---|---|---|---|
| **DF** | –6.8091 | –342.2664 | –344.2664 | –347.1106 | –345.4219 |
| **ADF(1)** | –4.3386 | –336.7035 | –339.7035 | –343.9698 | –341.4368 |
| **ADF(2)** | –3.8375 | –336.5126 | –340.5126 | –346.2009 | –342.8237 |
| **ADF(3)** | –3.3973 | –336.1947 | –341.1947 | –348.3051 | –344.0836 |
| **ADF(4)** | –3.1447 | –336.1198 | –342.1198 | –350.6524 | –345.5865 |

The Dickey-Fuller regressions include an intercept but not a trend.
Sample period: June 1991 to December 2001.
95% critical value for the augmented Dickey-Fuller statistic = –2.8842.

LL = Maximized log-likelihood
AIC = Akaike Information Criterion
SBC = Schwarz Bayesian Criterion
HQC = Hannan-Quinn Criterion

**Table 3. Inflation measurement in the UK.**

| **Dependent variable:** $\Delta\Delta NIFS_t$ | **(a)** **Jun 1996 to Dec 2001** | **(b)** **Jun 1996 to Mar 2003** |
|---|---|---|
| Constant | 0.27173 (0.30832) | 0.14557 (0.28015) |
| $\Delta\Delta RMed_t$ | 0.10237* (0.059474) | 0.094291 (0.053426) |
| $\Delta\Delta NIFS_{t-1}$ | –1.5325** (0.12546) | –1.4766** (0.11159) |
| $\Delta\Delta RMed_{t-1}$ | 0.054279 (0.083601) | 0.028796 (0.073860) |
| $\Delta NIFS_{t-2}$ | –1.7229** (0.21578) | –1.6672* (0.18618) |
| $\Delta RMed_{t-2}$ | 0.018022 (0.098925) | 0.0085138 (0.087837) |
| Bias | –15.0782 (94.9433) | –17.0982 (200.8351) |
| No. observations | 67 | 82 |
| $F$ test | $F(5,61) = 36.7646$** | $F(5,76) = 42.1724$** |
| Chow test[1] | $F(6,70) = 1.1595$ | — |
| Serial correlation[2] | $\chi^2(12) = 6.6030$ | $\chi^2(12) = 6.7978$ |
| Heteroskedasticity[3] | $F(1,65) = 0.28620$ | $F(1,80) = 0.10856$ |
| Residuals unit root test | $DF$=–7.8266* [–4.9838[4]] | $DF$=–8.6971* [–4.9290[4]] |

Notes:

Standard errors in brackets.

1. Chow's (1960) second test of equality between sets of coefficients. Sample period for structural stability: Jan 2002 to Mar 2003.
2. Lagrange Multiplier test of residual serial correlation (Godfrey, 1978a, 1978b).
3. Test based on the regression of squared residuals on squared fitted values (Koenker, 1981).
4. 95% critical value for the Dickey-Fuller statistic.

* Statistically significant at 5%.

** Statistically significant at 1%.

**Table 4. Inflation measurement in Italy.**

| **Dependent variable:** $\Delta\Delta NIFS_t$ | **Sample period:** **Mar 1991 to Dec 2001** |
|---|---|
| constant | 0.0058342<br>(0.12438) |
| $\Delta\Delta RMed_t$ | 0.059211<br>(0.037825) |
| $\Delta\Delta NIFS_{t-1}$ | –1.3410**<br>(0.089882) |
| $\Delta\Delta RMed_{t-1}$ | 0.037468<br>(0.042934) |
| $\Delta NIFS_{t-2}$ | –1.2195**<br>(0.14906) |
| $\Delta RMed_{t-2}$ | 0.034948<br>(0.038682) |
| Bias | –0.16694<br>(3.5565) |
| No. observations | 130 |
| $F$ test | $F(5,124) = 64.6541$** |
| Chow test[1] | $F(6,144) = 3.2044$** |
| Serial correlation[2] | $\chi^2(12) = 12.8727$ |
| Heteroskedasticity[3] | $F(1,128) = 0.13292$ |
| Residuals unit root test | $DF$=–11.1923* [–4.8425[4]] |



Notes:

Standard errors in brackets.

1. Chow's (1960) second test of equality between sets of coefficients. Forecast period: Jan 2002 to Feb 2004.
2. Lagrange Multiplier test of residual serial correlation (Godfrey, 1978a, 1978b).
3. Test based on the regression of squared residuals on squared fitted values (Koenker, 1981).
4. 95% critical value of the Dickey-Fuller statistic.

* Statistically significant at 5%.

** Statistically significant at 1%.

# Table 5. Inflation bias in Italy.

| Inflation bias (percentage points) | Chow test |
|---|---|
| 1% | 3.0582** (0.008) |
| 2% | 2.8855** (0.011) |
| 3% | 2.7001* (0.016) |
| 4% | 2.5159* (0.024) |
| 5% | 2.3438* (0.034) |
| 6% | 2.1906* (0.047) |
| 7% | 2.0596 (0.062) |
| 8% | 1.9512 (0.077) |
| 9% | 1.8636 (0.091) |
| 10% | 1.7944 (0.104) |



Notes:

Chow's (1960) second test of equality between sets of coefficients ($F(6,144)$).
Sample period for estimation: Mar 1991 to Dec 2001.
Sample period for structural stability: Jan 2002 to Feb 2004.
Significance levels in brackets.
* Statistically significant at 5%.
** Statistically significant at 1%.

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